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LOCAL LABOR MARKET CONDITIONS: EFFECTS ON LABOR DEMAND AND WAGE LABOR SUPPLY DECISIONS OF FARM AND RURAL NONFARM COUPLES, 1978-82*

by

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Local Labor Market Conditions: Effects on Labor Demand and Wage Labor Supply Decisions of Farm and Rural Nonfarm Couples, 1978-82

By J. G. Tokle and Wallace E. Huffman

The objective of this paper is to examine the effects of geographical differences in local economic conditions--both anticipated (permanent) and unanticipated (transitory)--on wage labor demand and labor force participation decisions of U.S. farm and rural nonfarm couples. Farm and rural nonfarm households in some regions of the U.S., e.g., the upper midwest, were adversely affected by changes in local economic conditions that occurred during 1978 to 1982 (U.S. Dept. Agr. 1987). Households in the southwest faired better. Part of these effects seems to be driven by changes in local labor market conditions.

This study distinguishes between anticipated and unanticipated components of economic variables because they are expected to have different effects in labor markets. One hypothesis is that localities that have higher anticipated employment growth and unemployment rates pay a wage premium to attract workers from other localities. Thus, when the local anticipated employment growth rate declines (relative to the national rate) real wage rates can be expected to fall in these localities and to rise in others. Another hypothesis is that localities experiencing unanticipated negative labor market shocks--a relative disturbance in labor demand or a higher unanticipated unemployment rate, for example as the upper midwest did during 1981-82--show a decline of real wage rates. These unanticipated negative disturbances to demand may be shared by workers and employers. Other geographical wage differences are undoubtedly due to cost of living and locational amenities.

Labor supply decisions consist of both a participation and hours of work decision by those who participate, and they seem to be distinct decisions

(Mroz 1987). In this study, we focus on the effects of geographical differences in local markets, primarily labor markets, on wage-work participation decisions. These effects are examined in a two-worker, risk-neutral household-decision-making framework rather than in a single-worker model.

A surprising amount of research on labor supply of married males and females has been in the context of a single-worker model (DaVanzo, DeTray, and Greenberg 1976; Mroz 1987, Table 1). Furthermore, Mroz (1987) has continued this tradition in a recent analysis of the sensitivity of female labor supply to an array of economic and statistical assumptions. We, however, believe that significant new insights into labor supply decisions and behavior of households can be gained by considering labor supply decisions in a two-worker husband-wife model. Huffman and Lange (1989) employ such an economic and econometric model in their recent study. In particular, a spouse's wage has substitution and income effects rather than just income effects. All earnings, work (income) conditioned transfers, and self-employment income are excluded from the other household income used to measure income effects. A spouses' education may affect a partners' labor supply decision by changing their reservation wage through efficiency effects in household production (and in the case of farm households in farm production) or tastes. A couple's labor supply decisions are affected by the same economic shocks, and taking these into account in a joint estimation procedure can be expected to increase the statistical efficiency of the parameter estimates.

The paper has the following organization. First, a picture of the empirical setting for the study is presented. Second, the economic models of labor demand and labor supply are developed with local labor market effects included. Third, the data and econometric model are summarized. Fourth, the econometric results are deployed and evaluated. Finally, some conclusions and implications are presented.

The Empirical Setting

To set the stage for the analysis, this section presents summary measures of differences across localities in labor market characteristics. The focus is on unemployment, employment growth, and shocks to labor demand. The basic geographic unit of analysis here and in the later econometric work is a state. States are the smallest geographic unit for which annual data on several labor market characteristics are readily available. Data are presented for 23 states that have the largest rural population in 1980, and they accounted for 76.5 percent of the total U.S. rural population.

Equilibrium unemployment rates differ geographically and do not tend to converge over time. Table 1 presents the ratio of predicted state unemployment rates to the predicted national unemployment rate for 1970, 1974, 1978, and 1982.[1/] Only four years are reported for each state to conserve on space for this slowly changing variable, and they lead up to and include the period 1978-82 which is the period covered in the later econometric analysis.

As has been pointed out in earlier studies (e.g., Hall 1972; Abowd and Ashenfelter 1981; Adams 1985; Topel 1986), at any point in time there are important regional differences in unemployment rates that exist and tend to persist over time. For example, the unemployment rates are uniformly higher than average for Michigan, West Virginia, Mississippi, and Louisiana. For Pennsylvania, Minnesota, Iowa, Virginia, North Carolina, Georgia, and Texas, the unemployment rates were uniformly below average. For other states, there is, however, not a simple relationship during 1970-1982 between their unemployment rate and the national average rate.

Employment growth rates differ geographically, and some of the differences persist during 1970-1982. Table 1 presents the ratio of predicted state employment growth rates to the predicted national employment growth rate.[2/] It shows that the employment growth rate in Georgia, Florida, Louisiana, Texas, and California were uniformly larger than average, and for New York, Pennsylvania, Ohio, Illinois, and Missouri, employment growth was uniformly below the national average. During this period, some states went from having far above average employment growth rates to having far below average employment growth. These states were Iowa, West Virginia, South Carolina, Kentucky, Tennessee, Alabama, and Mississippi. For some other states--Indiana, Michigan, Wisconsin, Minnesota, and North Carolina--the turn-around was only slightly less dramatic during this period. Thus, these data show a very unequal geographical distribution of employment growth during 1970-1982.

The geographical distribution of labor demand "shocks"--unanticipated changes in employment due to shifts in demand for output, the business cycle, weather, and other factors--is also unequal. Table 2 presents indexes of local labor demand shocks for 23 states for the subperiod 1978-1982, which is the period covered in the later econometric analysis. These measures were constructed as follows. First, the natural logarithm of annual private employment (1968-1982) was regressed on quadratic trend. The residuals from these regressions, ϵ_t^s, are indexes of time varying local demand conditions in state s and year t. Second, the natural logarithm of national aggregate employment was regressed on quadratic trend. The residuals from this regression, ϵ_t, are a measure of the aggregate labor demand disturbance in year t. Relative local labor demand disturbances in year t and state s are then defined as $\eta_t^s = \epsilon_t^s - \epsilon_t$, which expresses the current local labor demand "shock" as a deviation from the aggregate labor demand "shock." (This is a measure used by Topel 1986, pp. S129.)

These measures of local labor demand disturbances have two important features. First, for a given year there is substantial geographical variation in employment disturbances. They, also, seem to be as a group largely unrelated to the aggregate business cycle. Second, for a given state the successive disturbances tend to take the form of local cycles. Shocks with the same sign tend to persist for a couple of years then to reverse themselves. The upper midwestern states--Michigan, Indiana, Illinois, Wisconsin, Minnesota, and Iowa--were hit relatively hard with negative shocks during the recession that set in during 1981-1982. Some other states, for example, New York, Georgia, Florida, and Texas, experienced negative shocks during 1978-1979 and experienced positive relative labor demand shocks during 1981-1982.

This picture of considerable geographical variation of permanent and transitory characteristics of local labor markets can be expected to affect wage rates where there is considerable geographical immobility of workers and firms. Workers are largely immobile in the short run, but in the long run, they are geographically mobile. However, localities differ in their net advantages to firms (industries) and in their amenities and cost of living characteristics to workers. These seem to contribute to permanent differences in labor markets of different localities.

The Economic Model

A labor market transaction is a tied sale in the sense that the worker simultaneously sells the service of his/her labor and buys the attributes of his/her job including location. Employers, also, simultaneously buy the services and characteristics of workers and sell the attributes. Therefore, the theory of equalizing wage differentials is one that stresses both supply and demand for labor, and the market equilibrium performs a matching or sorting function of allocating or assigning specific workers to specific firms and locations (Rosen 1986). An acceptable match occurs when the preferred choices of an employer and an employee are mutually consistent.

The actual wage paid can be considered to be the sum of two conceptually distinct transactions, one for labor services and worker characteristics, and another for job and locational attributes (Rosen 1986). The wage paid by employers to induce workers to undertake undesirable tasks or work at an undesirable location takes the form of a wage premium--a negative price of the job. The observed distribution of wages is one that clears markets over all worker characteristics and job/locational attributes.

Localities are assumed to be heterogeneous over characteristics that matter for optimal firm-industry (employer) and household (worker) decisions on location and other activities. In the short-run, firms (employers) and households (workers) are geographically immobile. Workers are also occupationally immobile. In the long run, firms and workers are assumed to be geographically mobile, and workers are occupationally mobile. This mobility is, however, costly because it consumes resources that could be allocated to alternative uses. This environment is one of spatially related competitive labor markets that possess equilibrium wage differentials. Unequal transitory shocks to labor demand in different localities are an added source of geographical differences in wage rates.

The wage elasticity of aggregate labor demand for a locality is negative, but individuals face a perfectly elastic demand for their labor. Labor demand or wage rates at a given location and for a given sex are assumed to be a function of skill or human capital (ζ) and job or locational characteristics--permanent or anticipated labor market conditions (Ω), transitory or unanticipated labor market conditions (ω), local cost of living (ψ), and locational amenities (Δ). This relationship is summarized as:

(1) $W^j = W^j\ (\zeta^j,\ \Omega,\ \omega,\ \psi,\ \Delta),\ j =$ M(males), F(females).

By permanent or anticipated labor market conditions, we refer to important conditions that local firms and workers could be expected to know and to include in rational decision making. These variables include the anticipated or long-run unemployment rate, anticipated rate of job growth, and anticipated change in the occupational composition of employment. When anticipated unemployment harms workers (households) more than firms (employers), firms or localities that have higher anticipated or long run unemployment rates will have higher equilibrium wage rates than other firms or localities (Hall 1972; Abowd and Ashenfelter 1981; Adams 1985; Topel 1986). The higher wage rates are needed to compensate for higher probabilities of unemployment.

When households (workers) bear most of the cost of geographical and occupational mobility, geographic differences in the anticipated rate of employment growth and in change of occupational composition of employment are also the source of compensating geographical wage differentials (Lilien 1982; Adams 1985; Topel 1986). Positive wage differences are expected and required to provide the economic returns to households for geographical or occupational mobility.

Locational differences in the cost of living could be decomposed into differences due to prices of goods traded among locations and of nontraded goods and services (Tolley 1974; Kenny and Denslow 1980). With competitive markets, the prices of traded goods in two areas will differ only by transport costs. These differences seem likely to be an insignificant source of cost of living differences. Prices of nontraded goods and services can be expected to differ much more between two localities. The price of housing-plus-access is an example of a nontraded good whose cost of production differs among localities. Furthermore, housing costs are a significant component (14-15 percent) of workers' household expenditures. Thus, locational differences in the price of housing-plus-access can reasonably be expected to be a source of interlocality wage differences.

Areas differ in the quantity and quality of locational amenities that they provide. Normal climatic conditions, e.g., average January and July temperature, are characteristics of the local environment that households and firms can be expected to know, and they can be expected to matter to households (workers) and possibly to firms. Other studies (e.g., Israeli 1977; Hock and Drake 1974; Kenny and Denslow 1980) have found significant effects of climatic conditions on wage rates.

Labor Supply and Labor Force Participation

Households make labor supply and labor force participation decisions of their members. In the modeling that follows, the focus is on single-family, husband-wife households. In this analysis, we have simplified the decisions faced by the household to ones that can be modeled in a one-period static model. Households, however, are assumed to be risk neutral about uncertain outcomes including employment or unemployment. This means that the expected wage, the wage adjusted from the probability of employment, is used in resource allocation decisions.

An analysis of participation decisions of farm and rural nonfarm households is presented. The nonfarm households are assumed to have only wage and asset income; farm households have asset income, self-employment income from their farm business, and the possibility of off-farm wage income. Thus, the wage-labor participation decisions of farm households are more complex than for rural nonfarm households (Huffman and Lange 1989; Strauss 1986).

The economic decision-making framework of these households is summarized in the set of equations (1)-(4). Equation numbers that contain "n" refer specifically to nonfarm and an "a" refers specifically to farm households. Other equations refer to both types of households:

(1) $U = U(T_h^M, T_h^F, Y; \zeta^M, \zeta^F, \Delta, \tau)$

(2n) $\bar{T} = T_m^j + T_h^j, \; T_m^j \geq 0, \; j = M, F$

(2a) $\bar{T} = T_f^j + T_m^j + T_h^j, \; T_m^j \geq 0, \; j = M, F$

(3n) $(1 - u^M) W_o^M T_m^M + (1 - u^F) W_o^F T_m^F + V = P_y Y$

(3a) $(1 - u^M) W_o^M T_m^M + (1 - u^F) W_o^F T_m^F + P_Q Q - W_x X + V = P_y Y$

(4a) $Q = Q(T_f^M, T_f^F, X; \zeta^M, \zeta^F, \Delta)$.

Farm and nonfarm households are assumed to derive utility from the leisure time of the husband and wife (T_h^j) and from goods purchased in the market (Y). Household utility also depends on husband's and wife's human capital (ζ^j), local climate (Δ), and other household characteristics (τ), e.g., number of children in the household, commuting distance to service centers, which are not current choices.

Farm and nonfarm households receive an endowment of time each year ($\bar{T}$) for the husband and wife, which is treated as being heterogeneous. In nonfarm households, the time of each adult is assumed to be allocated between work for a wage (T_m^j) and leisure (T_h^j). In farm households, time is allocated among work on their own farm (T_f^j), work for a wage (off-farm) (T_m^j), and leisure

(T_h^j). In farm and nonfarm households, optimal hours of wage work might be zero in any year. Hence, a non-negativity constraint is imposed on wage work ($T_m^j \geq 0$).

The cash income of these farm and rural nonfarm households is uncertain. Although they may desire to supply T_m^j hours to the local market for wage work, they cannot be assured of being paid for all of these hours because of local unemployment. Thus, the wage rate, given employment, is adjusted for the expected probability of unemployment (u^j). These sex specific unemployment rates are assumed to be exogenously determined for the households. Furthermore, firms and households cannot be expected to know the unanticipated parts of employment growth or unemployment rates when they make plans. W_o^j is the anticipated wage, given unemployment, when unanticipated local labor disturbances are zero, i.e., $\omega = 0$ in equation (1). Expected household wage income is $(1 - u^M)\, W_o^M\, T_m^M + (1 - u^F)\, W_o^F\, T_m^F$. Household asset income is denoted as V. Farm households also have uncertain self-employment or net income from a farm business ($P_Q Q - W_x X$). The production of farm output (Q) is by inputs of husband's and wife's farm hours (T_f^j) and by purchased inputs (X). The efficiency of the production process is affected by human capital of the husband and wife (ζ^j) and climate (ϕ).[3/]

Households are assumed to face a perfectly elastic supply of the consumption good (Y) at a price P_y. Farm households, also, are assumed to face a perfectly elastic supply of inputs at a certain price W_X and perfectly elastic demand for farm output. The price of farm output is, however, uncertain when production plans are made, and P_Q denotes the expected price.

For nonfarm households, wage labor supply functions are obtained by maximizing (1) subject to (2n) and (3n). For farm households, wage labor supply functions are obtained by maximizing (1) subject to (2a), (3a), and (4a). These wage labor supply equations are:

$$(5n)\quad T_m^j = S_m^j\,[(1-u^j)W_o^j,\ (1-u^k)W_o^k,\ P_y,\ V,\ \zeta^M,\ \zeta^F,\ \Delta,\ \tau],\ T_m^k>0,$$

$$S_m^j\,[(1-u^j)W_o^j,\ P_y,\ V,\ \zeta^M,\ \zeta^F,\ \Delta,\ \tau],\ T_m^k = 0,$$

$$j,k = M,F;\ j \neq k.$$

$$(5a)\quad T_m^j = S_m\,[(1-u^j)W_o^j,\ (1-u^k)W^k,\ P_y,\ V,\ P_Q,\ W_X,\ \zeta^M,\ \zeta^F,\ \Delta,\ \tau,\ \phi],\ T_m^k>0;$$

$$S_m\,[(1-u^k)\,W_o^j,\ P_X,\ V,\ P_Q,\ W_X,\ \zeta^M,\ \zeta^F,\ \Delta,\ \tau,\ \phi],\ T_m^R = 0,$$

$$j,k = M,F;\ j \neq k.$$

Equations (5n) and (5a) show that wage labor supply functions for a given individual have different structures due to the outcome on their spouse's decision to work for a wage. (Also, see Huffman and Lange 1989.) When both married nonfarm individuals work for a wage, their wage labor supply is a function of their expected anticipated wage rates, price of consumption goods, asset income, their human capital stocks, the local climate, and tastes. For farm households, in addition the expected price of farm output, the price of purchased farm inputs, and a technology parameter determine off-farm wage labor supply.

Wage-labor participation decisions of an individual can be reduced to a comparison of his (her) reservation and anticipated market wage offers. The reservation wage equations are derived from the wage-labor supply equations by setting hours of wage work equal to zero and rearranging to express the wage, now the reservations wage, in terms of the other determinants of labor supply. Also, the determinants of the spouses wage are substituted for his (her) wage.

The reservation wage equations are:

$$(6n) \quad W_r^j = \left[\frac{1}{1-u^j} \right] G_r^j \, [P_y, V, \zeta^M, \zeta^F, \Delta, \tau, (1-u)^k, \Omega, \omega = 0, \psi],$$

$$j,k = M,F; \; j \neq k.$$

$$(6a) \quad W_r^j = \left[\frac{1}{1-u^j} \right] G_r^j \, [P_y, V, P_Q, W_X, \zeta^M, \zeta^F, \Delta, \tau, \phi, (1-u)^k,$$

$$\Omega, \omega = 0, \psi] \; j,k = M,F; \; j \neq k.$$

A nonfarm household member participates in wage work when his (her) reservation wage is less than the anticipated wage he (she) can expect to earn in the market.[4/] A farm household member participates in nonfarm wage work when the marginal value of his (her) leisure and (or) farm work hours are less than his (her) anticipated nonfarm wage.

The probability of wage work can be expressed as the probability that an individual's reservation wage is less than his (her) anticipated market wage. For the i-th household and j-th married individual, define

$$D_i^j = \begin{Bmatrix} 1 \text{ if j-th individual works for a wage} \\ 0 \text{ otherwise} \end{Bmatrix}, \; j = M,F,$$

then the probability of wage work for the i-th individual is:

$$(7) \quad P_r \, \{D_i^j = 1\} = F[W_{ri}^j < W_{oi}^j]; \; j = M,F, \text{ or}$$

$$(7a) \quad P_r \, \{D_i^j = 1\} = F[P_y, V, \zeta^M, \zeta^F, \Delta, \tau, \Omega, (1-u^j), (1-u^k)],$$

$$j,k = M,F; \; j \neq k, \text{ and}$$

$$(7n) \quad P_r \, \{D = 1\} = F[P_y, V, P_Q, W_X, \phi, \zeta^M, \zeta^F, \Delta, \tau, \Omega, (1-u^j),$$

$$(1-u^k)], \; j,k = M,F; \; j \neq k,$$

where F( ) is a distribution function. Variables that explain the probability of wage work are all of the variables that enter an individual's labor demand and labor supply functions, except for the individual's anticipated wage rate. When the labor supply schedule has a positive slope, variables that cause the labor supply curve to shift to the left will increase the reservation wage and reduce the probability of wage work. A change in a variable that raises the market wage--raises labor demand curve-- will increase the probability of wage work.

Selected variables are examined for their effects on the probability of wage work. An increase of an individual's schooling will increase his (her) anticipated market wage and reservation wage. The net effect on the probability of wage work is a priori ambiguous, but other studies (e.g., Heckman and MaCurdy 1980, 1982; Huffman and Lange 1989) have found a strong positive effect of an individual's schooling on his (her) probability of wage work for married farm and nonfarm males and females in the United States.

The effects of local labor market conditions depend partially upon whether they are anticipated or unanticipated. If the demand for leisure increases as the expected wage decreases, an increase of the expected local unemployment rate will raise an individuals' reservation wage. When anticipated unemployment hurts workers more than firms, firms or localities that have higher expected unemployment rates will pay higher wage rates. Thus, the net effect on the probability of wage work depends on which of these changes is largest. If a one percentage point increase in the expected unemployment rate results in a one percent increase of the anticipated wage, the probability of wage work will be unaffected. If the anticipated wage increases by less than one percent, then the probability of wage work will decrease.

Localities having more rapid anticipated job growth are expected to pay higher wage rates than other areas. The anticipated wage rates are higher in these localities, and the reservation wage seems likely to be unaffected. Thus, a higher expected growth rate of jobs will increase the probability of wage work.

On the other hand, unanticipated changes in local labor market conditions--relative shocks to employment growth and unanticipated unemployment--are not expected to affect the probability of wage work of married farm and nonfarm males and females. The reason is that these disturbances can at best be poorly forecasted and at worst cannot be forecasted.

When leisure is a normal good, an increase in the expected price of farm output will reduce the probability of wage work by couples who operate a farm business. The reason is that the quantity demanded of husband's and wife's farm labor increases provided their farm hours are a normal input and the quantity of their leisure demanded increases due to the increase of expected profit. Given the constraint on total time, the hours of wage work will decrease. The effects of farm input price changes on the prob ability of wage work are a priori ambiguous. Hired (nonfamily) labor and family labor seem likely to be heterogenous. If hired farm labor and husband's and wife's farm labor are gross substitutes, then an increase of the wage for hired labor can be expected to increase the demand for husband's and wife's farm labor. The increase of the input price reduces expected farm profit and reduces the demand for leisure of the husband and wife. Thus, the net effect on the reservation wage and probability of wage work is a priori ambiguous. A similar conclusion arises when one considers the effect of a change in the price of other inputs.

The Data and Econometric Model

The empirical analysis that follows focuses on the effects of local economic conditions on wage labor demand and probability of wage work of married rural males and females. By focusing on males and females, we can obtain a much better picture of the total effect on households, and on differences due to gender. The empirical analysis is based on farm and rural nonfarm households contained in the Current Population Surveys (CPS) and having economic activity in 1978, 1979, 1981, and 1982.

The Data

The CPS is a monthly survey containing information about the employment status of members of approximately 60,000 interviewed households residing in every state of the United States. The annual demographic file (March) contains information on labor force participation, employment, and earnings of household members during the calendar year before the survey. Starting in 1977, the state of residence and farm-nonfarm residence of each household is identified.

There are two samples collected for analysis. One sample consists of nonmetropolitan-nonfarm households where the husband and wife are present and no self-employment income was received. This is the closest definition to rural nonfarm that we can obtain because households are not identified in the CPS by rural nonfarm residence. This sample comprises what we label as the "rural nonfarm" wage earning households. The other sample consists of husband-wife households that have a farm residence and self-employed income from farming. In both samples, households having a residence in the contiguous 48 states, except for New England, are included. Households in New England were excluded because of the lack of importance of agriculture in those states. These two samples from the 42 states consist of about 8,115 rural nonfarm households and 1,466 farm households (per year).

Local markets are defined as state units in this study. The primary reason is that a state is the smaller geographic unit in which a CPS household can be identified. States are also the smallest political-economic-geographic unit for which annual data are collected on employment, unemployment, and agricultural prices. In addition, government programs frequently target state units. The major disadvantages are (i) state units are in some cases too large and heterogeneous to adequately summarize the economic conditions facing individual households and (ii) households may reside in one state but work and engage in most of their economic activity in an adjacent state. However, other recent studies (e.g., Adams 1985; Topel 1986) have used state units as labor markets.

Four years of economic activity during the period 1978-1982 are focused on to given variation in local economic conditions that affect wage and participation rates. This period was chosen for several reasons. First, we were constrained by 1977 being the first year that the CPS identified the farm-nonfarm residence of households. Second, the period 1975-79 is the trough-to-peak part of a national business-cycle expansion (Executive Office of the President 1987). The national average unemployment rate was 8.3 percent in 1975, and it declined to 5.8 percent in 1979. The late 1970s was also a period when net farm income was relatively good. The period starting in 1980 is one with a business-cycle contraction. The national unemployment rate rose from 7 percent in 1980 to 9.5 percent in 1982-83. The sharp rise of real interest rates and fall in the value of the U.S. dollar were contributing factors to the drop in net farm income during 1981 and 1982. Although the depression of the farm economy continued after 1982, extending the analysis through 1983 did not seem wise because 1983 is the year for the first large government payment-in-kind (PIK) programs. Third, the data in Table 2 show relatively large geographical variation in shocks to labor demand during this period. Fourth, Lilien (1982) has shown that the 1970s was a period when employment growth was very unequally distributed across U.S. industries. Significant shifts of the occupational-industrial mix of employment were occurring, especially a rise in the share of workers employment in services, and finance, insurance, and real estate, and decrease in share employment in manufacturing.

When the four cross-sectional files are combined together, the data on 42 states and 4 years gives 168 potentially distinct observations on different local economic conditions. This is considerable potential variation for helping to explain geographical differences in wage rates and participation in wage work. Although we have data for four years, it is not a panel consisting of the same households. About 25 percent of the CPS households in any year are replaced.

Empirical Definitions

Short empirical definitions and sample mean values of the variables are presented in Table 3. More detail is presented below on the derivation of selected variables.

Average hourly wage rates are derived for rural nonfarm married males and females. For individuals in the rural nonfarm wage work households, the average wage of an individual is his (her) wages and salaries for the year preceding the survey (1979, 1980, 1982, 1983 are the survey years) divided by the product of his (her) hours worked per week last year and weeks worked last year. For individuals who are in a household that has self-employment income, an accurate measure of the average wage rate cannot be computed. The reason is hours worked include hours at all jobs, both wage work and self employment.

Thus, average wage rates are not available for CPS farm household members. Nominal wage rates are deflated by the consumer price index to express them in real terms.

Five measures of local labor market conditions are derived. The predicted state employment growth rates ($PJOBGR_t$) is the indicator of anticipated change in local labor demand. It is obtained as the difference in forecasted values of the natural logarithm of a state's private sector employment in t and t-1. The forecasts were obtained from a regression of the natural logarithm of employment, 1968-1982, on a quadratic trend. We have chosen to measure a state's unemployment rate for all private sector employees rather than having separate rates for males and females. We believe that the combined rate is the best indicator of local conditions. The predicted state unemployment rate ($PURATE_t$) is the measure of the anticipated local unemployment rate. It is obtained from a regression of a state's annual unemployment rate, 1968-1982, on quadratic trend. The change in the share of a state's employment that is in the service sector is an indicator of change in the occupational mix of local labor demand. It is defined as the share in t minus the share in t-2. Service jobs include employment in services, transportation, government, finance and wholesale and retail trade. The derivation of the relative shocks to labor demand in state labor markets ($ESHOCK_t$) is described on page 4. In this study, these shocks are treated as being unanticipated by firms and workers. The residual or unanticipated unemployment rates ($RURATE_t$) is derived as the actual unemployment rate in t minus the predicted unemployment rate for t.

Geographical differences in the cost of living and locational amenities are tied to cost of housing-plus-access and climate. The price of land is one major part of the cost of housing-plus-access. For households living in rural areas, the base price of land is represented by the average agricultural land price in 1978 (U.S. Dept. Comm. 1980). The price of home sites is larger in areas where agricultural land prices are higher. If the land is expected to be converted to urban uses in the future, its price will be higher, and the cost of producing other local goods increases as the base land price increases (Kenney and Denslow 1980). In addition, the price of housing-plus-access increases as the percentage of the population living in urban areas increases because the cost of land plus commuting is larger in urban areas. The percentage of the population living in urban areas in 1980 (U.S. Dept. of Comm. 1981) is used as a second proxy for the cost of housing-plus-access.

The easiest locational amenities to measure are climate. We use 30-year (1950-80) normal average January and average July temperatures (Weiss, Whittington, Teigen 1985). Kenny and Denslow (1980) and Hoch and Drake (1974) found nonlinear effects of these temperatures on $\log_e$ wage rates in earlier studies.

The profitability of local agriculture is represented by indexes of agricultural prices and agricultural climate. Indexes of crop prices, livestock prices, wage for farm labor, and prices of other inputs are derived and deflated by the consumer price index. Two output price indexes are constructed because the average labor intensity of livestock production is significantly different for these output groups (Huffman and Evenson 1989, Ch. 10). The crop price index is composed of prices of 26 different commodities or commodity groups. The livestock price index is composed of 7 commodity groups. The expected prices that are used to derive the Fisher-type output price indexes are primarily one-year lagged prices of the commodities (Huffman and Evenson 1989, Ch. 10).

Input prices are split into two groups, farm labor and other inputs. The price of farm labor is the hourly wage paid to employees working for cash wages only. The nonlabor input prices include fertilizer, feed, capital, seed, land, and miscellaneous inputs. Current prices are used to derive Fisher-type price indexes of these inputs (Huffman and Evenson 1989, Ch. 10). Both output and input price indexes are deflated by the consumer price index to make them in real terms.

Normal annual precipitation and normal growing-degree-days are two important climatic variables for agricultural production that farmers can be expected to know when they make production plans. Normal annual rainfall is a 25 year average of annual precipitation (Weiss, Whittington, Teigen 1985). Natural precipitation is the primary source of water for much of U.S. agriculture. However, in low precipitation areas, irrigation is a costly substitute. GDD, accumulated growing degree days, is a measure of accumulated heat units from a temperature range that is particularly favorable to corn production (U.S. Dept. Agr. and U.S. Dept. Comm. 1970; U.S. Dept. of Comm. 1971, 1981). Corn is grown in almost every state, but more generally the index is highly correlated with good growing conditions for warm-season crops.

Household asset income is defined as household income from interest, dividends, and rental property deflated by the consumer price index. Thus, all income- and work-conditioned transfers are excluded.

The Econometric Model

The econometric model consists of two labor demand equations and two wage-participation equations. The empirical specification of the labor demand or wage equations, equation (1), is similar for married males and females:

$$(8)\quad \ln WAGE_i = \alpha_1^j + \alpha_2^j EXP_i + \alpha_3^j EXP_i^2 + \alpha_4^j ED_i + \alpha_5^j RACE_i + \alpha_6^j PJOBGR_i$$

$$+ \alpha_6^j RURATE_i + \alpha_7^j PURATE_i + \alpha_8^j \Delta SHRSER_i + \alpha_9^j ESHOCK_i + \alpha_{10}^j RURATE_i$$
$$+ \alpha_{11}^j \ln PLAND_i + \alpha_{12}^j URBAN_i + \alpha_{13}^j JAN_i + \alpha_{14}^j JAN_i^2 + \alpha_{15}^j JULY_i$$
$$+ \alpha_{16}^j JULY_i^2 + \alpha_{17}^j NC_i + \alpha_{18}^j SOUTH_i + \alpha_{19}^j WEST_i + \alpha_{20}^j TIME_i + \alpha_{21}^j \lambda_i^j$$
$$+ \epsilon_i^j, \; i = 1, \ldots, n, \; j = M,F.$$

The natural logarithm of an individual's real wage is expressed as a function of his (her) own human characteristics--experience, experience squared, education, race--and job/local conditions that are potential sources of geograph- ical wage differentials. The last group of variables includes sets of variables for local labor market conditions, cost of living and locational amenities, and regional dummy variables. A time trend and sample selectivity variables (λ_i^j) are also included in each equation.

A few additional details about the specification of equation (8) are highlighted. Experience, defined as an individual's age minus years of schooling completed minus 6, is included rather than an individual's age. This measure of experience is a reasonable proxy for useful work experience of males but less so for females because they spend a larger share of their lifetime in household activities. This measure of experience is less endogenous to current labor market decisions than actual experience (Heckman and MaCurdy 1980). When a sample selection variable is included in the wage equation, sample selection bias is unlikely to be associated with the use of a work experience variable (Mroz 1986).

Equation (8) is specified as being quadratic in experience, and January and July temperatures but not in other variables. This choice was made largely based on evidence reported in other studies, e.g., Adams 1985; Topel

1986; Kenney and Denslow 1980, and what worked well in preliminary fits of the wage equations. The regional dummy variables may contain redundant information about labor demand. If the set of variables representing local labor market conditions and cost of living and locational amenities proxy relatively well the sources of geographical wage differentials, then the regional dummy variables will not make a statistically significant contribution to real wage rates.

When women spend more time out of the labor force and are more geographically immobile (i.e., more likely to be a tied spouse) than men, women's wage rates can be expected to be less responsive to both their measured human capital and to local economic conditions (Mincer 1978). In these conditions, the coefficients or marginal effects of variables in the female wage equation will be smaller than in the male equation.

Sample selectivity and autocorrelation are potential problems in the wage equation. If a sample selection variable was not included in equation (8), the disturbance term of the wage equation would be expected to have a nonzero mean because the equation is fitted to a nonrandom subset of the total population (Heckman 1980). This is a potential source of statistical bias in estimated coefficients. With a selection variable included, the disturbance has a zero mean, and a normal distribution.

The data files consist of two adjacent year cross sections (1978-79, 1981-82) that are separated by one year. Some autocorrelation might be expected to exist in adjacent year observations on the same individual, but observations that are one or more years apart are less likely to be correlated. This short, disjointed time series where the composition of the sample changes over time is not conducive to being corrected for autocorrelation. Failure to correct for autocorrelation when it is present results in some loss of estimation efficiency, but the least-squares estimator is unbiased or consistent (Johnston 1984). Contemporaneous correlation of disturbances in the two wage equations might occur. However, there are unequal numbers of observations in the wage equations of the married males and females. This means that the two wage equations cannot be estimated jointly in the seemingly-unrelated regression model. Ignoring contemporaneous cross-equation correlation of disturbances results in some loss of estimation efficiency (Johnston 1984).

The empirical specification of equation (7a), the probability of the j-th married individual in the i-th farm household participating in wage work, is:

$$
\begin{aligned}
\text{(9a)}\quad Pr(D_i^j = 1) = F[&\beta_1^j + \beta_2^j AGEM_i + \beta_3^j AGEM_i^2 + \beta_4^j EDM_i + \beta_5^j EDF_i + \beta_6^j RACE_i \\
&+ \beta_7^j KIDS06 + \beta_8^j KIDS618 + \beta_9^j \ln ASSETINC_i + \beta_{10}^j PJOBGR_i + \beta_{11}^j PURATE_i \\
&+ \beta_{12}^j \Delta SHRSER_i + \beta_{13}^j ESHOCK_i + \beta_{14}^j RURATE_i + \beta_{15}^j \ln PLAND_i + \beta_{16}^j URBAN_i \\
&+ \beta_{17}^j JAN_i + \beta_{18}^j JAN_i^2 + \beta_{19}^j JULY_i + \beta_{20}^j JULY_i^2 + \beta_{21}^j \ln PCROP_i \\
&+ \beta_{22}^j \ln PLIVE_i + \beta_{23}^j \ln FARMWAGE_i + \beta_{24}^j \ln POTINP_i + \beta_{25}^j GDD_i \\
&+ \beta_{26}^j RAIN_i \cdot GDD_i + \beta_{27}^j NC_i + \beta_{28}^j SOUTH_i + \beta_{29}^j WEST_i + \beta_{30}^j TIME_i] \\
&j = M, F.
\end{aligned}
$$

where F( ) is the normal distribution function. This probability is a function of a set of variables representing an individual's own characteristics, his (her) spouse's characteristics, household characteristics, and local conditions that are a potential source of geographical differences in participation probabilities.[5/] This group of variables includes sets of variables for local labor market conditions, cost of living and locational amenities, agricultural prices and climate, and regional dummy variables. A time trend is also included.

A few additional details about equation (9) are highlighted. It is a reduced-form specification, including only variables that can be expected to be included in the labor supply or labor demand functions. It has a specification similar to the one used by Huffman and Lange (1989). The husband's age (and age squared) control for nonlinear life-cycle and work-experience effects when education is held constant. Age and our measure of experience are highly correlated, so experience is not included as a separate variable. A husband's and wife's ages are also highly correlated, and we employ only the husband's age for explaining both the husband's and wife's probability of wage work.

The participation equation is also quadratic in January and July temperatures. This is consistent with the specification of the labor demand equations. The effects of the agricultural climate on participation include an interaction variable between length of growing season (GDD) and annual precipitation (RAIN). If GDD and RAIN are substitutes in their effects on the reservation wage, the sign of this variable will be negative.

The participation equations for rural nonfarm household members are similar to those for farm household members, except the variables that refer specifically to the profitability of agriculture are excluded. These are the prices of agricultural outputs and inputs, annual precipitation, and the length of growing season.

Given that participation decisions of a husband and wife in farm and nonfarm households are assumed to be joint within a household optimizing framework, the probability of a given married individual participating in wage work is affected by some of his (her) spouse's characteristics. Also, these decisions are affected by random or unmeasured shocks to labor supply (reservation wage) and labor demand. These shocks are expected to be correlated across spouses (Huffman and Lange 1989). Thus, the estimation procedure for equation (9) will be by bivariate probit. If a husband and wife are generally affected similarly by a given shock, the correlation between these disturbances will be positive. For example, a favorable shock to labor demand will increase the probability of the husband and wife participating in wage work. If decision-making in the household is such that a husband and wife react differently to shocks in labor supply and labor demand, the correlation between the disturbances in the two participation equations will be negative.

The Results

Results from testing hypotheses about the effects of local economic conditions and other variables on labor demand for rural nonfarm married males and females and on the probability of wage labor participation of married males and females in farm and rural nonfarm households are reported and evaluated in this section.

Labor Demand

Labor demand equations are fitted to the data for 24,571 married rural nonfarm males and 17,508 married rural nonfarm females. Fitting of these equations (8) is by least squares with an instrumental variable included to control for sample selectivity. Although labor demand facing farm males and females is important, the available CPS data do not permit constructing a measure of the wage for farm household members. It is our belief, however, that conclusions reached about labor demand for married rural nonfarm males and females are applicable to married farm males and females, too.

Four estimated labor demand functions are reported in Table 4; two for males and two for females. They differ only in that the second equation excludes the regional dummy variables. A test of the null hypothesis that the coefficients of the three regional dummy variables are jointly equal to zero is rejected for males but not for females. The sample value of the F statistic is 5.08 for males and 1.90 for females, and the critical value of the F statistic with 3 and infinite degrees of freedom is 3.79 at the 1 percent significance level. Thus, for males the eleven variables representing local labor market conditions, cost of living differences, and location amenities do not capture all of the geographical differences in wage rates. However, for females the regional dummy variables are not adding significant information about geographical wage differences.

In the wage equations, all of the coefficients of the human characteristics have expected signs and are significantly different from zero at the 1 percent level. An increase of an individual's experience has first a positive but diminishing marginal effect on the real wage. The maximum effect of experience on $\log_e$ real wage occurs at 36.9 years for males and 31.5 years for females. This pattern has been reported in many studies. However, the $\log_e$ wage-experience relationship is more convex for males than for females; real wage rates rise faster for small amounts of experience for males than females, and after reaching a peak they also decline faster. This result is consistent with married females having on average less actual labor market experience for any given measured experience than for married males (Mincer and Ofek 1982). It is also consistent with males making larger investment in experience during their early work-life than females.

A one-year increase in schooling causes a larger percentage increase of the female than the male wage, 7.1 versus 5.5 percent. These relative magnitudes are consistent with some results reported in other studies, e.g., Topel

(1986) for males and Gerner and Zick (1983) for females. The average wage for married rural nonfarm working females in the CPS sample is 57.8 percent of the wage for males. Thus, the absolute or dollar value increase in the real wage for an additional year of schooling is larger for males than for females ($.16 versus $.12 per hour).

Nonwhite rural nonfarm males earn 20 percent less than rural nonfarm white males, other measured variables equal, and nonwhite females earn 6.5 percent less than white females. Topel (1986) found an 18 percent difference in wages of white and nonwhite males for 1976-79, which is similar to our estimate. Other studies have also shown large gaps in the wage of white and black men on average but little or no gap in white and black womens' wage rates (e.g., Hammermesh and Rees 1984). Thus, these results show relative differences in wage rates by race that are similar to other studies.

The set of variables representing local labor market conditions perform remarkably well. All of the signs of coefficients on these variables are positive, except for the coefficient of the unanticipated unemployment rate, in both the male and female wage equations, and they are all consistent with hypotheses developed earlier. Most of the coefficients of these variables in the male wage equation are significantly different from zero at the 5 percent level. For females, the marginal effects of the local labor market variables are smaller for variables than can be anticipated than for males but are as large or larger for variables that cannot be anticipated. Except for the unanticipated unemployment rate, the local labor market variables are statistically weaker in the female than in the male wage equation.

Strong evidence exists that the real wage rates of married males incorporate compensation for local labor market conditions that can be anticipated by employers and workers. The evidence is much weaker for females. This relative difference would be expected when married females spend a smaller share of their time in the labor force and are tied to their husband's locational choice (Mincer 1978). For married males, wage rates are higher in localities that have higher expected rates of employment growth and higher expected unemployment rates.

The wage premium in localities having higher rates of expected growth of labor demand is needed to provide compensation to males for the costs of geographical (and possibly occupation) mobility. Topel (1986) also found a similar effect of expected employment growth on real wage rates of nonfarm males. Our results show that a one percent increase of the expected growth rate of local employment increases the male real wage by 1.5 to 2.5 percent. For females, the marginal effect is about 1 percent. Thus, these results imply that married rural males and females in the upper midwest did experience significant reductions in the demand for their labor and real wage rates during the early 1980s when employment growth rates fell far below (about 2 percentage points) the national average in these states (see Table 1).

Localities that have higher anticipated unemployment rates also pay higher real wage rates for males. This result is similar to the findings of Abowd and Ashenfelter (1981), Adams (1985), and Topel (1986). When unemployment is anticipated and harms workers more than employers, localities (firms) that have higher expected unemployment rates pay a wage premium to compensate workers for bearing this predictable risk of unemployment. The evidence for compensating females is, however, much weaker. The directional effect is the same, but the statistical significance is quite low. Again, this finding is consistent with greater immobility of married females than males.

Our results imply the wage premium associated with higher anticipated unemployment rates is large enough for males to keep expected real wage rates approximately unaffected. A one percent increase of the expected unemployment rate causes a 1.1 to 1.2 percent increase of the male real wage rate. For females, this does not occur, and we have low confidence in the point estimate.

An increase locally of the share of service jobs increases the real wage. The effect is significantly different from zero at the 5 percent level for males but not for females. Although service jobs span a wide range of skills from motel and restaurant staff to investment bankers and rates of pay, compensation is needed to create the incentives for males to invest in skills and change occupations.

Real wage rates seem to respond to unanticipated changes in local labor market conditions in a way that is consistent with employers and employees sharing good and bad outcomes. Positive local relative labor demand disturbances increase real wage rates and negative shock decrease them. Topel (1986) found similar effects on wage rates of nonfarm males. The coefficients of ESHOCK are significantly different from zero at the 7 percent level (Table 4, equation 1) in the male wage equation and at the 20 percent level in the females' wage equation. The marginal effect on the real wage of a one percent relative shock to labor demand is, however, much smaller than a one percent change in expected employment growth or the expected unemployment rate. This seems to be a result of the fact that neither employers nor employees can accurately forecast these disturbances or plan for them.

Real wage rates seem to be flexible to local business-cycle downturns, especially for females. The coefficient of RURATE is significantly different from zero at the 10 percent level in the female wage equation but only at the 25 percent level in the male wage equation. This result would be reasonable if married males on average have significantly larger investment in firm-specific human capital than married females. Workers having firm-specific human capital are less likely to be laid-off (Becker 1975). This is a likely possibility given that females on average have shorter expected wage-work lives than males, and are a poorer place for firms to invest firm-specific human capital (Mincer and Ofek 1982).

Other studies have not looked at the effect of unanticipated unemployment on the female wage. Heckman and MaCurdy (1982) did find a negative effect of the current local unemployment rate on the female wage. Adams (1985), however, found a statistically significant positive effect of unanticipated unemployment rate on the male wage in his study. Thus, our results differ slightly from other studies for the effects of unanticipated unemployment on the female real wage.[6/]

As a group the variables representing geographical differences in cost of living and locational amenities contribute significantly to explaining wage rates of married rural nonfarm males and females. Under the null hypothesis that the six coefficients of these variables are jointly equal to zero, the sample value of the F statistic is 19.8 for males and 8.9 for females. The critical value of the F statistic with 6 and infinite degrees of freedom at the 1 percent significance level is 2.1. Thus, these null hypotheses of no effect are rejected. Wage rates of males and females differ because of local differences in cost of living and locational amenities.

The variables representing cost of living differences perform largely as expected. An increase of the land price increases significantly the wage rates of males and females. The point estimates of the elasticities are .060-.073 for males and .053-.056 for females. These magnitudes fall in the range predicted by Kenny and Denslow (1980) for wage adjustments to compensate for home site cost differences when housing costs are 15 percent of consumption expenditures and the price of the home site account for 21 percent of the value of a home, including site. An increase of the urban share of the population increases significantly the wage rate of males. URBAN, however, does not have a statistically significant effect on the wage of females.

Climate is a proxy for locational amenities, and the effects of JAN and JULY on wage rates is conditioned by whether the regional dummy variables are included. Because JAN and JULY are correlated with the regional dummy variables, the best measure of winter and summer temperatures on wage rates is obtained from the equations where the regional dummy variables are excluded (Table 4, equations (2) and (4)). The effect of JAN on $\log_e$ male wage is quadratic and statistically significant. An increase of JAN first increases the males wage, has a peak effect at 26°F, and then decreases the male wage. One might conclude that a normal average January temperature of 26° is associated with the worst of winter activities--too warm for snow and ice sports and too cold (and cloudy) for golf, tennis, and bicycling. Pennsylvania, Ohio, and Indiana have values of JAN close to 26 degrees. In the female wage equation, JAN and JAN^2 do not have statistically significant effects.

The effects of JULY on $\log_e$ wage is quadratic and statistically significant for males and females. For males, an increase of JULY first causes a reduction of the wage but at a decreasing rate until it reaches the low point at 82°F which is the maximum observed value. This type of relationship was also found by Hoch and Drake (1974). For females, the quadratic effect goes in the opposite direction. The peak occurs at 72°F which is near the bottom of the observed values and corresponds to average July temperatures in Pennsylvania and S. Dakota. Thus, for most of the observed values of JULY, an increase of JULY reduces the wage rates of males and females.

Participation in Wage Work

The bivariate probit estimates of the equations explaining the probability of wage work for farm and rural nonfarm couples are reported in Table 5. The equations were fitted to 32,662 observations on rural nonfarm households and 5,865 observations on farm households. Marginal effects of the regressors on the probability of wage work are evaluated at the sample mean and reported in Table 6.

The first two columns of Table 5 present results for rural nonfarm couples, and the last two columns are for farm couples. For both the farm and rural nonfarm couples, the estimated cross-equation correlation coefficient of the disturbances in the participation equations is positive--0.26 for farm couples and 0.19 for rural nonfarm couples, and they are significantly different from zero at the 1 percent level. This implies: (i) that the random disturbances in married male and female wage-work participation decisions are affected in the same direction by random shocks (or unmeasured effects); (ii) that the wage-work participation decisions of married males and females are not statistically independent. The choice functions should be fitted as one bivariate probit equation rather than two independent univariate probit equations.

Lets turn to specific results. For farm and nonfarm males and rural nonfarm females, the life-cycle effect on probability of wage work is quadratic. At young ages, an increase of an individual's age increases the probability of his (her) wage work. The maximum effect occurs at age 26.2 and 33.2 for farm and rural nonfarm males, respectively, and at age 20.8 for rural nonfarm females. At older ages, the probability of wage work decreases as the individual becomes older. For farm females, the probability of wage work is largest at a young age and decreases as they become older. The effects of AGEM on female participation may be dominated by cohort effects. Labor force participation rates of married women in younger cohorts are significantly higher than in older cohorts (Killingsworth and Heckman 1986).

A husband or wife who has more schooling has a higher probability of wage work. Thus, additional schooling raises an individual's market wage by more than it raises their reservation wage. These results for females are similar to ones reported by Heckman (1980) and Nakamura and Nakamura (1981) for non-farm married women. For farm males, these are similar to ones reported by Sumner (1982) and Huffman and Lange (1989). The marginal effect of a year of female schooling on her probability of wage work is larger than of male schooling on his probability of wage work (.41 vs .91 for nonfarm males and females and .15 vs .16 for farm males and females). The implication is that a year of schooling increases the difference between a wife's reservation and market wage relatively more than for her husband. This is consistent with results reported by Huffman and Lange (1989) for a different data set. Also, the marginal effect of schooling on the probability of wage work for a given gender is larger for rural nonfarm than for farm adults.

Although cross-person effects of education are seldom included (e.g., Heckman 1980, Nakamura and Nakamura 1981, Heckman and MaCurdy 1980, 1982, Sumner 1982), an increment to a spouse's schooling can change the probability of a mate participating in wage work. This occurs when the spouse's schooling causes the reservation wage of the mate to increase. Negative and statistically significant effects of a farm or rural nonfarm husband's schooling on his wife's participation and of a farm wife's schooling on her husband's wage work participation occurs. Huffman and Lange (1989) also found a similar negative effect of a farm wife's schooling on her husband's probability of wage work. For rural nonfarm males, additional schooling of their wife tends to increase the probability of their husband working for a wage.

Additional children at home ages less than six and 6-18 have well known and statistically significant negative effects on the probability of wage work by married females. The largest reduction occurs for additional children less than age six--about 5 percent per child. For older children, the negative marginal effect is larger for farm than for rural nonfarm married females. The coefficients of KIDS06 and KIDS618 are negative in both participation equations for married males, but they are generally weaker statistically. These results imply that additional children at home raise the reservation wage of married women relatively more than the reservation wage of married men. However, additional children ages 6-18 do cause a statistically significant reduction of the probability of wage work of farm males. The reduction for husbands is, however, about half the reduction for wives.

Although no other empirical study has examined the effects of local labor market conditions on the probability of wage work, the effects of these variables seem to be largely as expected. Anticipated variables have statistically stronger effects than unanticipated ones. A higher rate of local employment growth raises the market wage of males and females, and it increases the probability of wage work. These effects are significantly different from zero at the 5 percent level, except for farm wives. The results imply that a 2 percentage point drop of PJOBGR, e.g., as occurred in the upper midwest during the early 1980s, reduced the probability of wage work of rural nonfarm couples and farm wives by about 1 percent and of farm husbands by 2.2 percent. These are economically significant changes.

Recall that the results for the labor demand equations showed that a

1 percent change in the expected unemployment rate caused the male wage rate to rise by 1.1 to 1.2 percent but the female wage rate to rise by only 0.3 to 0.4 percent. Thus, if households are risk neutral, the effects of PURATE on participation are consistent. A larger PURATE increases the expected wage of males and increases the probability of wage work by then, but it reduces the expected wage of females and the probability of wage work of rural nonfarm females. There is no significant effect, however, on the probability of participation of farm females.

Although an increase of the share of service sector employment (ΔSHRSER) increased the wage rates of males, that change does not significantly affect the probability of wage work of males or females.

The relative employment shocks and unanticipated unemployment rate, which cannot be anticipated, do not have statistically significant effects on the probability of wage work of farm or rural nonfarm couples. Because they are unknown at planning time, no effect is expected.

Localities having higher land prices pay higher wage rates. Higher local land prices also increase the cost of housing which affects the reservation wage. If housing and husband's and wife's leisure are substitutes, then an increase of PLAND will increase the reservation wage rates. This makes the effect of PLAND on the probability of wage work <u>a priori</u> uncertain. However, if housing and husband's and wife's leisure are complements, their reservation wage rates will decline. The positive and statistically different from zero effect of PLAND on the probabilities of wage work implies that market wage effects are larger than reservation wage effects. A larger share of the population located in urban areas tends to reduce the probability of wage work of married males and females.

The effects of January and July temperatures on the probability of wage work are mixtures of effects through the market wage and reservation wage. Statistically significant effects on participation occur for nonfarm females and farm males. For farm males, a higher JAN first increases the probability of wage work; the peak effect occurs at 26°F; and then it decreases.

For farm households, the farm output and input price indexes have consistent signs, and their marginal effects are larger and are statistically stronger in the husband's participation equation. A drop of crop and livestock output prices increases the probability of wage work of husbands and wives. A decrease of the expected crop and livestock price indexes, by 5 percent, other things equal, would increase the husband's probability of wage work by 55 percent. A higher farm wage increases the probability of off-farm wage work of husbands. Other input prices have negative coefficients, but they are not significantly different from zero.

Agricultural climatic effects are important for farm males. Larger normal annual rainfall or a longer growing season increases the probability of wage work. RAIN and GDD are substitutes in their effects on the probability of wage work.

The coefficients of the regional dummy variables provide estimates of broad regional effects that are not captured in other regressors. Except for women living in the south, the regional effects are not significantly different from zero. Married women living in the south, however, have a lower probability of wage work than women in the northeast. Other things equal, the reduction is about 1 percent for rural nonfarm women and 5.5 percent for farm women.

The probability of wage work for rural nonfarm women has a positive and statistically significant trend. None of the other coefficients of TIME are significantly different from zero. In particular, there is not a similar positive effect of TIME for farm women.

Conclusions and Implications

This study has presented an econometric analysis of the effects of geographical differences in local economic conditions on labor demand and labor force participation. The fitted labor demand functions show that wage premiums exist for localities anticipating labor demand growth, higher unemployment rates, larger shares of employment in services, and higher costs of living. Some of the differentials seem, however, to be larger and in other cases smaller than what is required to keep utility constant for a risk neutral household. Unanticipated negative disturbances in local labor markets associated with employment growth and unemployment reduce wage rates.

Male wage rates are more responsive to anticipated changes in local labor market conditions, and female wage rates are more responsive to unanticipated changes. The differences by gender for anticipated effects are a result of females spending less time in the labor force on average than males. Women's wage rates bear an unusually large burden of negative unanticipated labor market outcomes because they are more geographically and occupationally immobile than men and are tied to locational decisions dominated by male employment considerations.

The main policy recommendations are that (i) public information be made available to firms and households about anticipated changes in all local labor markets, (ii) regionally targeted stabilization policies be employed to moderate unanticipated changes in local labor demand, and (iii) tax credits or other incentives be given to workers to hasten their movement from slow growing regions and occupations to rapidly growing ones.

Labor force participation decisions of households were shown to be affected by changes in anticipated local economic conditions when the difference between the expected market wage and the reservation wage became smaller or larger. For farm households, the probability of wage work increases when expected farm output prices decline and decreases when there is an expected decline in the growth of local labor demand. Thus, nonfarm earnings and farm income represent diversified income sources of farm households. Given the general trend toward larger farms, the results show that exiting from agriculture is significantly easier for farm households that reside in localities where nonfarm employment is growing rapidly than in other areas. These effects pull more strongly on males than females. Participation by farm and nonfarm couples is, however, unaffected by unanticipated labor market outcomes.

Future research could be undertaken to extend this analysis (i) through the late 1980s when national economic conditions and local labor fluctuations were moderate and (ii) to examine the effects of local economic conditions on hours of wage work through effects on wage rates.

Footnotes

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1/ Each state's annual unemployment rate, 1968-1982, and the national unemployment rate were regressed on quadratic trend. The respective predicted unemployment rates are forecasts from these regressions.

2/ The natural logarithm of total private sector employment, 1968-1982, was regressed on quadratic trend. Predicted employment growth rates are the first differences of the predicted employment values from these regressions.

3/ The effect of a tax on income is excluded from the economic (and econometric) model. This simplification is unlikely to be of major consequence for the empirical results (Mroz 1987).

4/ Alternatively, decisions can be made by comparing indirect utility functions associated with different outcomes.

5/ Although the number of children at home and the amount of household asset income is endogenous in a lifetime planning horizon, Mroz (1987) found that these variables were not endogenous in his tests. They are exogenous in our study.

6/ The null hypothesis that the 5 coefficients of the local labor market variables in the labor demand equation for females are jointly equal to zero cannot be rejected at the 1 percent significance level. For equation (3), Table 4, the sample value of the F statistic is 1.14, and the critical value for 5 and infinite degrees of freedom is 3.02 at the 1 percent significance level.

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Table 1. Geographical Distribution of Unemployment and Employment, 23 States, 1970-1982

States	Unemployment[a/] 1970	1974	1978	1982	Employment[b/] 1970	1974	1978	1982	Rural population 1980 (1000)
New York	1.09	1.25	1.17	.94	-.32	-.04	.22	.46	2,700
Pennsylvania	.89	.98	1.08	.90	.45	.35	.26	.17	3,643
Ohio	.89	.92	1.07	1.27	.86	.58	.30	.25	2,879
Indiana	.87	.85	1.04	1.33	1.09	.71	.30	-.04	1,965
Illinois	.83	.82	.97	1.20	.41	.40	.39	.38	1,908
Michigan	1.28	1.26	1.42	1.63	1.09	.66	.26	-.13	2,711
Wisconsin	.85	.74	.85	1.10	1.23	1.02	.82	.63	1,685
Minnesota	.83	.74	.75	.78	1.41	1.28	1.12	.97	1,351
Iowa	.63	.51	.64	.94	1.64	1.06	.52	-.00	1,206
Missouri	.87	.79	.86	1.00	.77	.71	.65	.59	1,567
Virginia	.72	.74	.78	.80	2.09	1.68	1.29	.92	1,817
W. Virginia	1.35	1.05	1.15	1.44	1.32	.97	.60	.29	1,244
N. Carolina	.80	.80	.85	.90	1.59	1.33	1.08	.84	3,059
S. Carolina	1.02	.92	.97	1.11	2.18	1.64	1.08	.55	1,433
Georgia	.83	.92	.92	.84	1.59	1.42	1.29	1.13	2,054
Florida	.85	1.08	1.06	.83	2.45	2.17	1.94	1.72	1,533
Kentucky	.96	.79	.89	1.15	2.09	1.42	.73	.13	1,799
Tennessee	.87	.80	.96	1.23	1.82	1.28	.78	.2	1,818
Oklahoma	.98	.87	1.10	1.52	1.73	1.33	.95	.59	1,556
Mississippi	2.11	1.57	1.25	1.00	2.09	1.64	1.08	.55	1,328
Louisiana	1.30	1.11	1.06	1.05	1.32	1.55	1.77	2.02	1,319
Texas	.78	.75	.74	.70	1.68	1.90	2.16	2.39	2,896
California	1.35	1.28	1.15	.93	1.23	1.37	1.51	1.60	2,060
United States	4.6	6.1	7.2	8.1	2.20	2.26	2.32	2.38	59,495

[a/]Numbers are ratios of the predicted state unemployment rate to the predicted national unemployment rate. Annual unemployment rates for each state and for the U.S. were regressed on quadratic trend, 1968-82, to obtain the predictions.

[b/]Numbers are ratios of the predicted state employment growth rate to the predicted national employment growth rate. The logarithm of annual employment in the private sector for each state and the U.S. were regressed on quadratic trend, 1968-82. The predicted employment growth is the difference between the predicted logarithm of employment in two adjacent years.

Table 2. Relative Employment Disturbance, 23 States, 1978-82[a]

States	1978	1979	1980	1981	1982
New York	-2.25	-2.24	-0.97	0.69	2.94
Pennsylvania	-0.40	0.35	-0.31	0.25	0.32
Ohio	0.67	0.98	-0.34	-0.13	-0.31
Indiana	2.58	2.22	-1.21	-0.49	-1.28
Illinois	0.90	0.71	0.81	-1.05	-0.90
Michigan	4.12	3.05	-0.57	-1.25	-2.16
Wisconsin	0.60	1.42	0.33	-0.62	-0.77
Minnesota	0.66	1.48	1.00	-0.28	-1.56
Iowa	1.89	0.95	0.11	-0.53	-1.87
Missouri	1.22	1.46	-0.37	-0.98	-0.37
Virginia	-0.52	-1.41	0.49	-0.15	0.34
W. Virginia	0.41	1.98	0.62	-1.40	-1.23
N. Carolina	-0.17	0.42	0.12	0.3	-0.17
S. Carolina	-0.72	0.67	0.06	0.57	0.47
Georgia	-0.43	-0.80	-0.49	0.11	1.65
Florida	-3.33	-2.85	0.14	1.88	2.55
Kentucky	1.01	1.29	-0.86	-1.13	-0.01
Tennessee	0.87	0.44	-0.77	0.39	-0.18
Oklahoma	1.79	0.43	-0.12	-1.00	-0.56
Mississippi	1.95	1.49	0.26	-0.97	-1.34
Louisiana	0.92	-1.03	0.23	0.40	-1.29
Texas	-1.09	-1.42	-0.72	0.76	0.68
California	0.67	0.85	0.84	0.12	-1.31
United States	2.39	3.59	1.88	0.34	-3.77

[a] The numbers are measures of $\eta_t^s = \epsilon_t^s - \epsilon_t$. See the text for details.

Table 3. Variable Names and Sample Means - Rural Married Couples 1978-79, 1981-82

Symbol	Variable description	Sample mean: Nonfarm	Sample mean: Farm
Individual/household			
AGEM	Husband's age (yrs)	47.0	50.5
AGEF	Wife's age (yrs)	43.9	47.2
EDM	Husband's schooling (yrs)	11.5	11.3
EDF	Wife's schooling (yrs)	11.6	11.8
RACE	1 if nonwhite; 0 otherwise	.07	.03
KIDS06	Number of children under age 6	.31	.27
KIDS618	Number of children ages 6-18	.66	.69
ASSETINC	Household real nonwage and nonfarm income (1967 prices) a/	$10,386 a/	$10,701 a/
Local labor market conditions			
PJOBGR	Predicted state employment growth rate (see text)	2.17	2.06
PURATE	Predicted state unemployment rate (see text)	7.46	7.04
ΔSHRSER	Change of share of a state's jobs in serv. occupation (prev.2yrs)	1.02	.91
ESHOCK	Relative state employment growth shock (see text)	.09	.15
RURATE	Residual state unemployment rate (see text)	-.37	-.34
Cost of living and locational amenities			
PLAND	State ave. price of agri. land in 1978 ($1,000/acres)	.39 a/	.79 a/
URBAN	Percentage of state population urban	.68	.66
JAN	Normal January ave. temperature (degree F)	34.1	29.7
JULY	Normal July ave. temperature (degree F)	75.9	75.1
Agricultural prices and climate			
PCROP	State real price index for crops (1967 prices)	-	.444 a/
PLIVE	State real price index for livestock (1967 prices)	-	.517 a/
FARMWAGE	State real wage rate for hired farm labor (1967 prices)	-	.504 a/
POTINP	State real price index for nonlabor farm input (1967 prices)	-	.536 a/
RAIN	State annual average precipitation (inches)	-	35.7
GDD	State ave. growing season length - growing degree days	-	3,336
Regional dummies and trend			
NC	1 for resident in North Central Region; 0 otherwise	.46	.28
SOUTH	1 for resident in South; 0 otherwise	.38	.51
WEST	1 for resident in West; 0 otherwise	.12	.06
TIME	Trend	3.00	3.00
Dependent variables			
WAGEM	Real married male nonfarm wage ($/hr, 1967 prices)	$2.87 a/	-
WAGEF	Real married female nonfarm wage ($/hr, 1967 prices)	$1.66 a/	-
D^M	1 if husband works for wage; 0 otherwise	.75	.43
D^F	1 if wife works for a wage; 0 otherwise	.54	.39

a/ Geometric mean. All other numbers are arithmetic means.

Table 4. Labor Demand Equations: Rural Nonfarm Married Males and Females, 1978-79, 1981-82

Variables	ln Wage			
	Males		Females	
	(1)	(2)	(3)	(4)
Human capital				
EXP (AGE-ED-6)	.031 (20.04)	.031 (19.95)	.017 (12.88)	.017 (12.93)
$EXP^2/100$	-.042 (10.30)	-.041 (10.20)	-.027 (9.35)	-.027 (9.38)
ED	.055 (40.21)	.055 (40.12)	.071 (30.12)	.071 (30.19)
RACE	-.204 (13.72)	-.203 (13.69)	-.065 (3.48)	-.064 (3.43)
Labor market conditions				
PJOBGR	.016 (2.68)	.024 (4.89)	.009 (1.19)	.011 (1.73)
PURATE	.012 (4.21)	.011 (4.13)	.004 (1.16)	.003 (0.76)
ΔSHRSER	.005 (1.91)	.005 (1.97)	.002 (0.72)	.002 (0.67)
ESHOCK	.005 (1.79)	.004 (1.45)	.005 (1.31)	.004 (1.07)
RURATE	-.006 (1.15)	-.004 (0.74)	-.011 (1.68)	-.010 (1.51)
Cost of living and locational amenities				
ln PLAND	.073 (5.46)	.060 (5.47)	.053 (3.02)	.056 (3.87)
URBAN	.255 (5.77)	.180 (5.33)	.011 (0.20)	.012 (0.26)
JAN	.003 (1.15)	.008 (3.96)	-.002 (0.55)	.002 (0.74)
$JAN^2/100$	-.001 (3.91)	-.014 (5.29)	.002 (0.57)	-.001 (0.32)
JULY	-.087 (1.43)	-.059 (1.25)	.197 (2.41)	.166 (2.59)
$JULY^2/100$	.057 (1.41)	.036 (1.14)	-.136 (2.49)	-.117 (2.77)
Regional dummies and trend				
NC	-.023 (1.37)		-.038 (1.67)	
SOUTH	.057 (2.65)		-.015 (0.51)	
WEST	.064 (2.09)		-.002 (0.05)	
TIME	-.026 (6.29)	-.027 (6.42)	-.012 (2.22)	-.013 (2.27
$\hat{\lambda}$	.279 (5.53)	.286 (5.68)	-.026 (0.79)	-.021 (0.82)
Intercept	6.070 (3.05)	2.260 (1.27)	-7.625 (2.50)	-6.404 (2.66)
R^2	.1619	.1614	.0781	.0780
N	24,571	24,571	17,508	17,508

Table 5. Bivariate Probit Estimates of Wage Labor Participation Equation for U.S. Farm and Rural Nonfarm Married Couples, 1978-79, 1981-82

	Wage work			
	Rural nonfarm		Farm	
Variables	Husband	Wife	Husband	Wife
Individual/household				
AGEM	.105	.027	.033	.003
	(26.82)	(7.92)	(3.78)	(0.30)
$AGEM^2/100$	-.158	-.068	-.063	-.036
	(41.07)	(19.69)	(7.28)	(3.83)
EDM	.044	-.016	.010	-.010
	(11.70)	(5.43)	(1.31)	(1.38)
EDF	.007	.095	-.029	.079
	(1.71)	(25.76)	(3.31)	(9.05)
RACE	.117	.333	.350	.403
	(3.55)	(11.20)	(3.23)	(3.75)
KIDS06	-.026	-.497	-.028	-.386
	(1.31)	(35.55)	(0.87)	(11.41)
KIDS618	-.008	-.084	-.037	-.072
	(0.83)	(10.99)	(2.23)	(4.33)
ln ASSETINC	-.338	-.829	-.347	-.542
	(4.26)	(10.57)	(2.68)	(4.08)
Local labor market conditions				
PJOBGR	.041	.052	.075	.036
	(2.87)	(4.50)	(3.31)	(1.60)
PURATE	.017	-.024	.050	.003
	(2.55)	(4.25)	(2.25)	(0.37)
ΔSHRSER	-.001	.003	-.020	-.007
	(0.21)	(0.55)	(1.23)	(0.40)
ESHOCK	-.005	-.001	-.014	-.017
	(0.71)	(0.16)	(0.91)	(1.12)
RURATE	-.013	-.012	-.004	-.032
	(1.07)	(1.18)	(0.13)	(1.03)
Cost of living and locational amenities				
ln PLAND	.168	.040	-.066	.027
	(5.97)	(1.72)	(1.03)	(0.41)
URBAN	-.078	-.190	-.188	-.094
	(0.79)	(2.47)	(0.62)	(0.31)
JAN	.007	-.020	.053	.022
	(1.05)	(3.99)	(4.42)	(1.80)
$JAN^2/100$	-.032	.019	-.077	-.010
	(4.35)	(3.08)	(3.59)	(0.45)
JULY	-.115	.498	.081	-.027
	(0.84)	(4.55)	(0.45)	(0.15)
$JULY^2/100$	.082	-.321	-.091	.014
	(0.90)	(4.38)	(0.75)	(0.12)
Agricultural prices and climate				
ln PCROP			-.251	-.135
			(1.33)	(0.71)
ln PLIVE			-.488	-.054
			(1.78)	(0.19)
ln FARMWAGE			.890	-.291
			(3.04)	(0.98)

Table 5. (Continued)

Variables	Wage work: Rural nonfarm Husband	Wage work: Rural nonfarm Wife	Wage work: Farm Husband	Wage work: Farm Wife
ln POTINP			-.520	-.589
			(0.84)	(0.93)
RAIN			.014	.004
			(1.19)	(0.30)
GDD/1,000			.323	-.014
			(2.82)	(0.12)
RAIN x GDD/1,000			-.038	-.013
			(1.61)	(0.55)
Regional dummies and trend				
NC	.069	.004	-.102	-.156
	(1.51)	(0.12)	(0.74)	(1.15)
SOUTH	-.019	-.115	-.038	-.314
	(0.39)	(2.93)	(0.24)	(2.00)
WEST	.007	-.047	-.287	-.247
	(0.10)	(0.80)	(1.61)	(1.40)
TIME	-.008	.018	.000	.007
	(0.77)	(2.09)	(0.02)	(0.28)
Intercept	6.702	-11.190	-.030	5.058
	(1.30)	(2.70)	(0.00)	(0.75)
Cross equation correlation coeff.	.187		.262	
	(15.32)		(12.16)	
Sample size	32,662		5,866	

Table 6. Marginal Effects (Percentage Point Changes) on the Probability of Wage Work

Regressors	Rural nonfarm		Farm	
	Husband	Wife	Husband	Wife
Individual/household				
AGEM	-.42	-.36	-.46	-.49
EDM	.41	-.16	.15	-.15
EDF	.07	.91	-.43	1.16
RACE	1.12	3.20	5.16	5.93
KIDS04	-.25	-4.77	-.41	-5.70
KIDS618	-.08	-.81	-.55	-1.06
ln ASSETINC	-3.25	-7.96	-5.11	-7.98
Labor market conditions				
PJOBGR	.40	.50	1.10	.53
PURATE	-.17	-.24	.73	.12
ΔSHRSER	-.01	.03	-.30	-.10
ESHOCK	.05	-.01	-.20	-.25
RURATE	-.13	-.11	-.06	-.47
Cost of living and locational amenities				
ln PLAND	1.61	.38	-.98	.39
URBAN	-.75	-1.82	-2.77	-1.39
JAN	-.15	-.07	.11	.24
JULY	.10	-9.46	-.81	-.09
Agricultural prices and climate				
ln PCROP	-	-	-3.70	-1.99
ln PLIVE	-	-	-7.19	-.80
ln FARMWAGE	-	-	13.11	-4.28
ln POTINP	-	-	-7.67	-8.69
RAIN	-	-	.21	.05
GDD	-	-	4.76	-.21
Regional dummies and trend				
NC	-.67	.04	-1.51	-2.30
SOUTH	-.18	-1.11	.56	-4.63
WEST	.07	-.45	-4.23	-3.64
TIME	-.08	.17	.01	.10